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Abstract: Menu prices from 13 international restaurant franchises that operate in both El Paso and Ciudad Juárez are utilized to examine the behavior over time of the peso/dollar exchange rate. Parametric and non-parametric tests indicate that the price ratio alone provides a biased estimator for the exchange rate. In addition to the multi-product price ratio, the empirical analysis also incorporates interest rate parity and balance of payment variables. The combination of unique microeconomic sample data with national macroeconomic variables illustrates one manner in which border economies provide information regarding the interplay of financial markets between Mexico and the United States.

Resumen: Precios de menú de restaurantes de 13 franquicias internacionales que operan en El Paso y Ciudad Juárez se utilizan para examinar el comportamiento del tipo de cambio del peso. Pruebas paramétricas y no-paramétricas indican que la relación de precios representa una medida sesgada para el tipo de cambio en términos estrictamente aritméticas. Además de la relación de precios de productos múltiples, el análisis empírico también incorpora tasas de interés de los dos países y una variable que refleja cambios en la balanza de pagos en México. La combinación de datos microeconómicos con datos macroeconómicos ilustra una de las maneras en que economías fronterizas ofrecen información acerca de los nexos financieros entre México y Estados Unidos.

Key Words: Prices, Exchange Rate, Border Economics

JEL Category: F31, Exchange Rates

Introduction

Exchange rate studies that analyze geographic and commodity group data have become relatively common in recent years (Engel and Rogers, 1996; 2001). This is in part due to the popularity of "hamburger" currency indexes that involve widely consumed and popular franchise restaurant menu items (Ong, 1997; Anonymous, 2002). It is also because this approach utilizes microeconomic data that were previously not available and complement the traditional macroeconomic data sets that rely upon aggregate price variables (Pakko and Pollard, 1996; Evans and Lyons, 2002).

This paper utilizes cross-border menu price data for milkshakes, pizzas, steaks, and other items to examine the exchange rate behavior of the Mexican Peso. The analysis takes advantage of a multi-component international restaurant price data set that matches menu items found in the sister cities of El Paso, Texas in the United States and Ciudad Juarez, Chihuahua in Mexico. On its own, the cross-border restaurant price index has been found to provide a biased estimate of the exchange rate between the peso and the dollar (Fullerton and Coronado, 2001). Accordingly, the modeling strategy employed herein goes beyond that implied by purchasing power parity (Balassa, 1964) to include elements suggested by interest rate parity (IRP) and balance of payment hypotheses (Aliber, 1973; Blanco and Garber, 1986; Throop, 1993; Zhou and Mahdavi, 1996).

Section 2 provides a brief overview of related exchange rate studies. The studies reviewed focus primarily on variants of purchasing power and interest rate parity models. The purchasing power parity (PPP) papers also include recent efforts that investigate international food price ratio comparisons to exchange rates. Section 3 discusses data collection efforts and the theoretical models utilized to analyze the data. A geographically unique price ratio is utilized as the PPP component. It is based upon a sample of approximately 70 menu prices whose values are collected at a monthly frequency. Interest rate information is calculated using 91-day Certificados de Tesoreria (CETES) in Mexico and 90-day Treasury Bills (T-BILL) in the United States. Balance of payment information is introduced by employing a ratio of international reserves to imports in Mexico. Empirical results are summarized in Section 4. Concluding remarks and suggestions for future research are discussed in the final section.

Literature Review

In an era of variable and occasionally volatile exchange rates, exchange rate monitoring has come to occupy a central role in both corporate planning and public policy analysis. Associated with the efforts to reduce foreign exchange rate risk, there is a great deal of interest in studies dealing with the determination of exchange rates. A large percentage of these studies rely upon PPP and IRP modeling frameworks (Marston, 1997). Short-run departures from PPP are fairly common, but a variety of papers report long-run evidence that favors different versions of this hypothesis (Jorion and Sweeney, 1996; Wu and Wu, 2001). Given its regional history of periodic financial instability, efforts to model Latin American exchange rates frequently rely on balance of payment information to augment the more widely-used PPP and IRP frameworks (Blanco and Garber, 1986; Fullerton, Hattori, and Calderón, 2001).

Several well-known techniques are based on PPP frameworks. One common approach is to calculate trade-weighted real exchange rate indexes. Under that method, an index number greater than 100 indicates overvaluation and an index number below 100 points to currency undervaluation (Fuentes, 2002). A second popular technique deals with implied nominal exchange rate calculations based upon national price index movements relative to a specific base period (Cheung and Wong, 2000; Lara y Beltrán del Río, 2002). A third approach was introduced more than a decade ago by *The Economist* magazine. "Burgernomics" takes advantage of an international food franchise menu to develop a simplified PPP index (Ong, 1997). The strategy relies upon using a globally popular hamburger as the homogeneous comparison good. The product selected is a signature menu item produced in 120 separate countries. While not intended to replace careful currency market analysis or more technically sophisticated monitoring devices, the burger index correctly signaled that the Euro would decline relative to the U.S. dollar following the introduction of the new currency in 1999 (Anonymous, 2002).

The popularity of the easy to understand hamburger index served as a catalyst for additional empirical efforts that make more extensive use of restaurant pricing patterns within the PPP framework. Pakko and Pollard (1996) examine the reliability of PPP and hamburger indexes for 15 currencies relative to the United States dollar. Results obtained indicate that the conditions for absolute PPP, under which ratios of national price indexes approximate exchange rates, do not hold in the short-run. Similarly, the relative version of PPP, which states that percent changes in the prices levels will lead to similar proportional changes for exchange rates, does not hold in the short-run. Potential explanations for those outcomes include barriers to trade, especially in agricultural products, causing prices of goods to differ across borders; variations in non-tradable good prices such as real estate and utilities leading to generalized price differences among countries; oligopoly market structures contributing to further price misalignments between regions; current account imbalances; and productivity gaps that also contribute to international price divergences. Those factors notwithstanding, Ong (1997) obtains results that indicate that a burger-based PPP index does hold in the long-run.

Other authors have utilized larger baskets of goods and consumer price sub-indexes to examine evidence implied by a cross section of products (Fraser, Taylor, and Webster, 1991; Engel and Rogers, 1996, 2001; Jenkins, 1997). This branch of the literature highlights several factors that can cause pricing patterns to deviate across markets. In particular, distances and transportation costs are generally found to contribute directly to the magnitudes of price differences between regions (Chen and Finney, 2002). Those deviations are also found to be greater in cases where international borders also serve to intensify market segmentation effects normally observed for metropolitan markets separated by distance. Regional business cycle and industrial composition differences have also been identified as sources of temporary price pattern and exchange rate divergences (Clark, Sawyer, and Sprinkle, 1997, 1999, 2001).

Additional factors can also affect price ratio comparisons between markets with different stages of development. A variety of papers (Balassa, 1964; Summers and Heston, 1991; Heston and Summers; 1996) argue that exchange rate conversions will overstate income estimates for higher income countries such as the United States and understate income estimates for lower income nations such as Mexico. Kakkar (2001) reports evidence that nontradeable price differentials play an important role in PPP deviations observed for the peso/dollar exchange rate. Also contributing to those numeric gaps are differing capital-labor ratios, menu costs, and taxes (Bhagwati, 1984; Rogers and Jenkins, 1995). As pointed out by Dornbusch (1976), currency market overshooting can also result from interest rate disparities combined with asset market and goods market adjustment differences. Given the above, it would not be surprising for price ratios for such country pairs to differ from exchange rates.

Much of the evidence reported by Fullerton and Coronado (2001) corroborates the potential divergence between currency quotes and restaurant price ratios for Mexico and the United States. The latter study examines menu prices between franchise restaurants in El Paso, Texas and Ciudad Juárez, México. On average, they sample more than 72 menu prices from approximately 13 distinct franchise restaurants found on both sides of the border (the numbers vary in response to menu changes, closures, and openings). In nearly two-thirds of the monthly observations, menu prices on the south side of the border are found to be cheaper than what is indicated by comparing them to the peso/dollar exchange rate and their companion products in the United States. The restaurant price ratio is found to be correlated with the exchange rate, but provide a biased predictor for it. Menu prices on both sides of the border change frequently in Fullerton and Coronado (2001), but not in a manner that appears related to variations in peso/dollar quotes. Some portion of the deviation between the price ratio and the exchange rate in that study may also result from the local-currency pricing effect identified for Mexico by Engel (2001).

Of course, PPP models represent only one approach to the analysis of currency markets. A large number of papers have reported at least partial evidence in favor of IRP modeling frameworks (Aliber, 1973; Gregory, 1987; Marston, 1997). Several of these efforts have been carried out with respect to the Mexican peso. In particular, Khor and Rojas-Suarez (1991) report empirical results that support the uncovered IRP hypothesis. The latter paper also highlights difficulties that the government in Mexico is likely to face if it attempts to lower interest rates without first attaining overall economic stability. Historically, financial disequilibria have played prominent roles in currency market volatility affecting the peso (Gil-Díaz and Carstens, 1996).

A variety of authors have examined factors that can cause developing country exchange rates to depreciate rapidly. Those efforts generally incorporate aspects of both PPP and IRP modeling strategies. A frequent approach employed for Latin American currencies includes balance of payment variables in the various model specifications (Blanco and Garber, 1986; Fullerton, Hattori, and Calderón, 2001). This paper relies upon a similar framework that utilizes a more extensive sample of the Fullerton and Coronado (2001) restaurant price ratio as the PPP component in the empirical analysis.

Data and Methodology

An elementary representation of one model that predicts movements in the peso by utilizing purchasing power and interest rate components is as follows:

$$NEX_t = \beta_0 + \beta_1 PR_t + \beta_2 IR_t \tag{1}$$

where NEX_t is the nominal exchange rate between the peso and dollar in month t, PR_t is the price ratio between restaurant products in Mexico and the United States in month t, and IR_t is the interest rate ratio between Mexico and the United States for the same period. Variants of this basic formulation have been employed in several earlier studies (Zhou and Mahdavi, 1996; Marston, 1997; Fullerton and Coronado, 2001).

The purchasing power parity component is calculated using the monthly ratio of menu prices from Ciudad Juárez and El Paso. Menu price data are obtained from a variety of cross border franchise operations. The latter include four hamburger chains, three pizza franchises, two fried chicken restaurants, two Mexican food establishments, one sandwich chain, and one upscale family restaurant. Two of the thirteen companies are headquartered in Mexico, the rest are from the United States. Data are collected by visiting the franchise pairs during the third week of each month in both cities. The restaurant list can vary whenever closures or openings occur on either side of the border. Additionally, individual food items have sometimes been dropped from the sample due to their elimination from menus in either or both cities. The sample period covered is July 1997 through June 2001.

The comparison items whose prices are sampled every month are largely homogeneous. Because they are food items designed to be eaten quickly, these meal items are not tradable goods in the classical sense. In fact, it is illegal to bring some of their contents across the bridges into the United States. The latter include items such as pork products, fresh fruit, and fresh vegetables that are used in many of the products included in the monthly sample. Although they are not tradable commodities, arbitrage opportunities exist in Ciudad Juárez where prices are quoted in pesos, but payments are accepted in either currency. Separate evidence reported for Mexico using national data series indicates that deviations between the exchange rate and the price ratio are likely due to nontradability (Kakkar, 2001). For the multi-product restaurant price ratio discussed herein, arbitrage pressures potentially serve to minimize the magnitudes and durations of any price inequalities that result from currency market shocks (Asplund and Friberg, 2001).

After collecting the raw data on both sides of the border, individual price ratios are then calculated by dividing the price in pesos by the price in dollars for each menu item. Statistical moments are also calculated for all of the monthly samples. The first means and variances are used to conduct t-tests for sample mean and exchange rate equality. Recent PPP studies (Fullerton and Coronado, 2001; Wu and Wu, 2001) indicate that monthly data utilized in these samples may not follow a normal distribution. Given that, the third and fourth moments, skewness and kurtosis, are used to conduct chi-square tests for sample distribution normality (Bera and Jarque, 1981). When non-normal sample data are encountered, a nonparametric test is used to test for sample mean and exchange rate equality. The procedure employed, a Wilcoxon signed-rank test, is distribution free (Daniel, 1978). In cases where the exchange rate and price ratio series are equal, the β_2 regression coefficient estimated for the interest rate ratio in Equation (1) will likely be statistically indistinguishable from zero.

For the interest rate variable shown in Equation 1, the ratio of 91-day Certificados de Tesorería (91-day CETES) yields from Mexico to 90-day United State Treasury Bill (90-day T-Bill) rates of return is employed. Data for the southern interest rate are available from the Banco de México web site (www.banxico.org.mx). For the northern interest rate variable, data are available from the Federal Reserve Bank of St. Louis website (www.stlouisfed.org).

Introduction of the import coverage ratio changes the basic model specification to that shown below:

$$NEX_t = \_{}_{0} + \_{}_{1}PR_t + \_{}_{2}IR_t + \_{}_{3}ICR_t \qquad (2)$$

Equation (2) reflects the role that balance of payment pressures occasionally play in Latin American currency markets (Blanco and Garber, 1986). To take that possibility into account, an import coverage ratio is employed. The latter variable is calculated as the ratio of monthly international reserves in Mexico, net of gold deposits, to monthly imports of goods and services measured in dollars. Data for both components of the import coverage ratio may be accessed via the central bank web site in Mexico (www.banxico.org.mx) or via other electronic media (International Monetary Fund, 2002). The exchange rate is hypothesized to vary inversely with respect to the latter variable, implying that $\_{}_{3}$ will be less than zero.

Empirical Results

Parametric and nonparametric statistical tests are conducted for equality between the average monthly price ratios and exchange rates. For the parametric approach, a standard t-test is conducted for equality between the arithmetic means of the price ratios and the average nominal exchange rate for every month of the sample. The 48 month sample includes four episodes of relatively rapid nominal depreciation during which the peso declined by 4-percent or more in a single month. Those periods include November 1997, October 1998, May 1999, and June 2000. The peso also appreciated relative to the dollar in a notable manner during several periods, including March 1999, July 2000, and April 2001 (Table 1).

Results for the t-tests appear in Table 2. In 26 of the 48 months for which sample data are available, the price ratio differs significantly from the corresponding exchange rate. In those 26 months the price ratio always falls below the exchange rates. This implies that menu prices in Mexico are less expensive than their counterparts in the United States, potentially reflecting labor cost differentials and other variables such as own-price and cross price elasticities of demand between the two economies (Heston and Summers, 1996). From January 1999 to June 2001 the price ratio does not differ significantly from the exchange rates in 22 of the 30 periods. While the apparent convergence of the two series is fairly impressive, the t-test employed assumes data normality. In order to examine whether the monthly bi-national restaurant data meet this requirement, a chi-square test is utilized (Bera and Jarque, 1981; Pindyck and Rubinfield, 1998).

Table 3 reports the results of the chi-square test for normality of the monthly price data. In all but seven time periods, the null hypothesis of the price ratio normality is rejected at the five-percent significance level. Consequently, a Wilcoxon signed-rank nonparametric test is introduced to minimize the risk of incorrect inference associated with the t-test results. The latter procedure is a distribution-free test and does not require assumptions regarding the density function of the variable examined (Daniel, 1978).

Results for the Wilcoxon signed-ranks test appear in Table 4. At the five-percent significance level, the null hypothesis of mean price ratio equality with the average nominal exchange rate is rejected in 35 of the 48 months for which sample data are available. As in the case of the t-test results shown in Table 2, all of the time periods in which the Wilcoxon signed-ranks tests fail to reject the null hypothesis occur in 1999, 2000, and 2001. That may imply that the initial evidence (Fullerton and Coronado, 2001) of exchange rate deviations from the price ratio PPP measure represent only temporary departures from the norm for the borderplex restaurant markets. If the two series are statistically equal to each other, it raises a question regarding shock dissipation of or speed of realignment when currency shocks occur.

As numerous authors have pointed out (Zhou and Mahdavi, 1996; Marston, 1997), such temporary deviations from PPP may result from interest rate differentials between trading partners such as Mexico and the United States. Deviations from PPP may also result from balance of payment fluctuations (Blanco and Garber, 1986). Table 5 contains regression output generated for an exchange rate equation that includes contemporaneous lags of the PPP restaurant price ratio, interest rate differential, and import coverage ratio variables. The results indicate that the coefficients for each of the regressors are statistically significant, but the coefficient for the interest rate ratio is greater than zero. This implies that if interest rates in Mexico rise relative to those in the United States, the peso will depreciate. That result runs counter to the hypothesized sign for _$_2$ as discussed above. The equation in Table 5 also includes a statistically significant autoregressive parameter at lag 1 to correct for serial correlation.

To confirm the results shown in Table 5, 26 separate versions of Equation 2 involving up to four lags of the explanatory variables were also estimated. In 25 out of 26 regressions, the interest rate ratio coefficient was both positive and significant. While counterintuitive, those outcomes confirm a positive relationship between the interest rate ratio and the exchange rate during the sample period studied. Several seminar participants and paper discussants have suggested that such a result may reflect monetary circumstances in Mexico wherein upward interest rate moves reflect the inflationary consequences of peso depreciations. The latter possibility may warrant investigation once additional sample data become available (for additional arguments along these lines, see Banco de México, 1998).

Given the apparently robust results associated with Table 5, several observations can be made with respect to the behavior of the currency market in Mexico. One is that the fairly strong linkage between the restaurant price ratio and the exchange rate reported by Fullerton and Coronado (2001) is confirmed. A second observation is that balance of payment shocks and other market developments contribute to periodic deviations from PPP-implied values of the peso. Perhaps not surprisingly, the information reported in Table 5 indicates that successful monitoring of the exchange rate in Mexico will involve a combination of PPP and other international finance measures. The borderplex restaurant price ratio seemingly offers one means by which this objective may be partially achieved.

The evidence to date indicates that the borderplex restaurant price index may help in predicting variations in the peso/dollar exchange rate. Because it does not involve traded goods, that possibility may seem surprising (Xu, 2003). It is important to remember, however, that arbitrage opportunities exist for border region restaurant customers since both currencies can generally be utilized for payment (Asplund and Friberg, 2001; Yoskowitz and Pisani, 2002). Because the data herein are exclusive to the borderplex, it would be helpful to assemble similar information for other points along the border between the two countries. Such efforts would eventually allow panel estimates to be employed in terms of examining deviations from parity and adjustment speeds to such shocks (Fleissig and Strauss, 2000).

Conclusion

A variety of research efforts in recent years have utilized multi-national franchise restaurant price comparisons with the objective of better understanding international currency valuations. Mixed evidence has been reported with respect to different versions of the purchasing power parity hypotheses tested. This paper extends one of those earlier efforts that indicated that a basket of cross-border menu prices provides a biased estimator for the peso/dollar exchange rate between Mexico and the United States.

In addition to taking advantage of a larger 48-month sample, the analysis also incorporates currency modeling strategies involving variables designed to reflect interest rate differentials between the two economies and balance of payment shocks that periodically affect Mexico. Estimation results confirm statistically significant relationships between the peso/dollar exchange rate and each of the three explanatory variables, but the sign for the interest rate differential variable is opposite of what is hypothesized. The signs of the restaurant price ratio and import coverage ratio parameters are as hypothesized. The estimation outcomes indicate that a strict interpretation of the PPP model is not supported for this short-run data. Nominal price differentials, as measured by the cross-border franchise samples, between Mexico and the United States do play important roles in monthly variations of the exchange rate.

Additional sampling will eventually permit a more complete set of tests to be performed. At present, the sample is not large enough to take advantage of a variety of time series techniques such as error correction specifications that permit disentangling both short- and long-term factors that potentially affect this currency market. New sampling will also allow tests to be conducted with respect to the length of time required for price deviations to dissipate following movements in the peso/dollar exchange rate. Additional tests could also be performed if similar data are collected for other cities along the border between Mexico and the United States.

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Table 1. Monthly Data Set (July 1997 – December 1999)

Month	Exchange Rate, P/$	Price Ratio	Mexico Interest Rates (91-(Día Cetes)	U.S. Interest Rates (90-Day T-Bills)	Interest Rate Ratio
Jul-97	7.89	6.64	19.40	5.05	3.84
Aug-97	7.79	6.35	20.15	5.14	3.92
Sep-97	7.79	6.32	20.51	4.95	4.14
Oct-97	7.88	6.37	19.91	4.97	4.01
Nov-97	8.26	6.65	22.01	5.14	4.28
Dec-97	8.15	6.68	19.88	5.16	3.85
Jan-98	8.25	6.89	19.37	5.04	3.84
Feb-98	8.49	7.00	19.63	5.09	3.86
Mar-98	8.62	6.88	20.76	5.03	4.13
Apr-98	8.50	6.87	19.47	4.95	3.93
May-98	8.61	6.77	18.85	5.00	3.77
Jun-98	8.91	7.13	20.99	4.98	4.21
Jul-98	8.90	7.06	21.82	4.96	4.40
Aug-98	9.96	7.43	25.22	4.90	5.15
Sep-98	10.11	7.67	41.90	4.61	9.09
Oct-98	10.15	7.84	37.53	3.96	9.48
Nov-98	9.94	8.04	34.30	4.41	7.78
Dec-98	9.87	8.50	34.35	4.39	7.82
Jan-99	10.17	8.81	32.27	4.34	7.44
Feb-99	9.94	9.15	28.72	4.44	6.47
Mar-99	9.52	9.41	23.86	4.44	5.37
Apr-99	9.29	9.34	21.05	4.29	4.91
May-99	9.75	8.98	21.02	4.50	4.67
Jun-99	9.49	9.08	21.35	4.57	4.67
Jul-99	9.38	9.01	20.78	4.55	4.57
Aug-99	9.37	8.90	21.49	4.72	4.55
Sep-99	9.36	9.05	21.34	4.68	4.56
Oct-99	9.65	9.19	20.30	4.86	4.18
Nov-99	9.36	9.10	18.68	5.07	3.68
Dec-99	9.51	9.19	17.65	5.20	3.39

Table 1. Monthly Data Set (January 2000 – June 2001)

Month	Exchange Rate, P/$	Price Ratio	Mexico Interest Rates (91-Día Cetes)	U.S. Interest Rates (90-Day T-Bills)	Interest Rate Ratio
Jan-00	9.48	9.10	17.43	5.32	3.28
Feb-00	9.41	9.24	16.44	5.55	2.96
Mar-00	9.29	9.04	14.46	5.69	2.54
Apr-00	9.40	9.01	14.37	5.66	2.54
May-00	9.52	8.74	15.58	5.79	2.69
Jun-00	9.96	8.84	16.61	5.69	2.92
Jul-00	9.36	8.91	14.62	5.96	2.45
Aug-00	9.23	9.00	15.71	6.09	2.58
Sep-00	9.41	9.24	16.15	6.00	2.69
Oct-00	9.64	9.38	17.06	6.11	2.79
Nov-00	9.41	9.57	18.01	6.17	2.92
Dec-00	9.57	9.51	17.41	5.77	3.02
Jan-01	9.67	9.30	18.50	5.15	3.59
Feb-01	9.70	8.94	18.07	4.88	3.70
Mar-01	9.62	8.73	16.47	4.42	3.73
Apr-01	9.35	8.93	15.40	3.87	3.98
May-01	9.10	9.10	12.61	3.62	3.48
Jun-01	9.15	9.28	10.27	3.49	2.94

Table 1. Monthly Data Set (July 1997 – December 1999)

Month	Mex. Int. Reserves (US$ Mil.)	Mexico Imports (US$ Bil.)	Import Coverage Ratio
Jul-97	24566		4.01
Aug-97	25841		4.22
Sep-97	26966		4.40
Oct-97	28102		4.59
Nov-97	27001		4.41
Dec-97	28797	73475.00	4.70
Jan-98	29186		4.23
Feb-98	29047		4.21
Mar-98	30118		4.36
Apr-98	31139		4.51
May-98	30968		4.49
Jun-98	30645		4.44
Jul-98	31679		4.59
Aug-98	29774		4.31
Sep-98	29266		4.24
Oct-98	30675		4.44
Nov-98	29766		4.31
Dec-98	31799	82816.30	4.61
Jan-99	31681		4.15
Feb-99	31494		4.12
Mar-99	31284		4.10
Apr-99	31470		4.12
May-99	31146		4.08
Jun-99	31346		4.10
Jul-99	32060		4.20
Aug-99	32067		4.20
Sep-99	32585		4.27
Oct-99	32268		4.22
Nov-99	31650		4.14
Dec-99	31782	91654.50	4.16

Table 1. Monthly Data Set (January 2000 – June 2001)

Month	Mex. Int. Reserves (US$ Mil.)	Mexico Imports (US$ Bil.)	Import Coverage Ratio
Jan-00	33643		3.58
Feb-00	33312		3.55
Mar-00	36371		3.87
Apr-00	34685		3.69
May-00	33566		3.57
Jun-00	32974		3.51
Jul-00	34323		3.65
Aug-00	32882		3.50
Sep-00	34108		3.63
Oct-00	35271		3.75
Nov-00	34690		3.69
Dec-00	35509	112735.04	3.78
Jan-01	39421		4.23
Feb-01	39106		4.20
Mar-01	40234		4.32
Apr-01	40309		4.33
May-01	40561		4.35
Jun-01	40759	111833.16	4.37

Table 2. Gossett t-Test for Price Ratio/Exchange Rate Equality (July 1997 – Dec. 1999)

Month	Sample Size	Computed t-statistic	Critical Value	Decision
Jul-97	75	-5.58	1.667	Reject
Aug-97	75	-7.513	1.667	Reject
Sep-97	74	-8.211	1.667	Reject
Oct-97	73	-7.501	1.667	Reject
Nov-97	73	-7.685	1.667	Reject
Dec-97	73	-6.863	1.667	Reject
Jan-98	73	-6.528	1.667	Reject
Feb-98	73	-6.21	1.667	Reject
Mar-98	73	-7.217	1.667	Reject
Apr-98	73	-6.761	1.667	Reject
May-98	73	-10.078	1.667	Reject
Jun-98	72	-6.165	1.667	Reject
Jul-98	72	-6.939	1.667	Reject
Aug-98	72	-8.225	1.667	Reject
Sep-98	72	-8.437	1.667	Reject
Oct-98	72	-7.167	1.667	Reject
Nov-98	72	-6.324	1.667	Reject
Dec-98	72	-4.065	1.667	Reject
Jan-99	72	-3.711	1.667	Reject
Feb-99	72	-1.943	1.667	Reject
Mar-99	72	-0.234	1.667	Fail to Reject
Apr-99	72	0.106	1.667	Fail to Reject
May-99	90	-2.28	1.667	Reject
Jun-99	94	-1.186	1.667	Fail to Reject
Jul-99	94	-1.118	1.667	Fail to Reject
Aug-99	94	-1.515	1.667	Fail to Reject
Sep-99	94	-1.073	1.667	Fail to Reject
Oct-99	82	-1.452	1.667	Fail to Reject
Nov-99	82	-0.805	1.667	Fail to Reject
Dec-99	82	-1.036	1.667	Fail to Reject

Table 2. Gossett t-Test for Price Ratio/Exchange Rate Equality (January 2000- June 2001)

Month	Sample Size	Computed t-statistic	Critical Value	Decision
Jan-00	81	-1.376	1.667	Fail to Reject
Feb-00	81	-0.403	1.667	Fail to Reject
Mar-00	81	-0.671	1.667	Fail to Reject
Apr-00	78	-1.180	1.667	Fail to Reject
May-00	81	-2.678	1.667	Reject
Jun-00	81	-3.970	1.667	Reject
Jul-00	81	-1.658	1.667	Fail to Reject
Aug-00	81	-0.867	1.667	Fail to Reject
Sep-00	81	-0.550	1.667	Fail to Reject
Oct-00	81	-0.808	1.667	Fail to Reject
Nov-00	81	0.5470	1.667	Fail to Reject
Dec-00	81	-0.195	1.667	Fail to Reject
Jan-01	74	-1.065	1.667	Fail to Reject
Feb-01	79	-3.288	1.667	Reject
Mar-01	86	-4.111	1.667	Reject
Apr-01	86	-1.886	1.667	Reject
May-01	86	-0.176	1.667	Fail to Reject
Jun-01	86	0.676	1.667	Fail to Reject

Table 3. Jarque-Bera Chi-Square Test for Price Sample Normality (July 1997 – Dec. 1999)

Month	Sample Size	Computed JB-statistic	Critical Value	Decision
Jul-97	75	1.898	5.991	Fail to Reject
Aug-97	75	5.294	5.991	Fail to Reject
Sep-97	74	1.929	5.991	Fail to Reject
Oct-97	73	27.664	5.991	Reject
Nov-97	73	16.616	5.991	Reject
Dec-97	73	15.054	5.991	Reject
Jan-98	73	18.865	5.991	Reject
Feb-98	73	119.388	5.991	Reject
Mar-98	73	126.558	5.991	Reject
Apr-98	73	127.693	5.991	Reject
May-98	73	1.925	5.991	Fail to Reject
Jun-98	72	99.076	5.991	Reject
Jul-98	72	141.138	5.991	Reject
Aug-98	72	110.016	5.991	Reject
Sep-98	72	20.697	5.991	Reject
Oct-98	72	23.488	5.991	Reject
Nov-98	72	42.416	5.991	Reject
Dec-98	72	39.504	5.991	Reject
Jan-99	72	28.809	5.991	Reject
Feb-99	72	21.816	5.991	Reject
Mar-99	72	28.621	5.991	Reject
Apr-99	72	11.797	5.991	Reject
May-99	90	86.871	5.991	Reject
Jun-99	94	154.861	5.991	Reject
Jul-99	94	171.613	5.991	Reject
Aug-99	94	111.094	5.991	Reject
Sep-99	94	82.809	5.991	Reject
Oct-99	82	72.74	5.991	Reject
Nov-99	82	89.043	5.991	Reject
Dec-99	82	76.326	5.991	Reject

Table 3. Jarque-Bera Chi-Square Test for Price Sample Normality (Jan. 2000 – June 2001)

Month	Sample Size	Computed JB-statistic	Critical Value	Decision
Jan-00	81	61.022	5.991	Reject
Feb-00	81	82.553	5.991	Reject
Mar-00	81	62.74	5.991	Reject
Apr-00	78	42.553	5.991	Reject
May-00	81	18.535	5.991	Reject
Jun-00	81	30.839	5.991	Reject
Jul-00	81	15.759	5.991	Reject
Aug-00	81	19.538	5.991	Reject
Sep-00	81	53.113	5.991	Reject
Oct-00	81	198.66	5.991	Reject
Nov-00	81	289.65	5.991	Reject
Dec-00	81	280.525	5.991	Reject
Jan-01	74	640.92	5.991	Reject
Feb-01	79	4.538	5.991	Fail to Reject
Mar-01	86	5.058	5.991	Fail to Reject
Apr-01	86	3.159	5.991	Fail to Reject
May-01	86	182.531	5.991	Reject
Jun-01	86	148.417	5.991	Reject

Table 4. Wilcoxon Signed-Rank Test, Price Ratio/Exchange Rate Equality (Jul97 – Dec99)

Month	N	T-	T+	T-*	T+*	CV	Decision
Jul-97	75	2325.0	525.0	4.753	-4.753	1.96	Reject
Aug-97	75	2530.0	320.0	5.835	-5.835	1.96	Reject
Sep-97	74	2510.0	265.0	6.047	-6.047	1.96	Reject
Oct-97	73	2426.5	274.5	5.916	-5.916	1.96	Reject
Nov-97	73	2439.0	262.0	5.984	-5.984	1.96	Reject
Dec-97	73	2384.0	317.0	5.682	-5.682	1.96	Reject
Jan-98	73	2376.0	325.0	5.638	-5.638	1.96	Reject
Feb-98	73	2378.0	323.0	5.649	-5.649	1.96	Reject
Mar-98	73	2439.0	262.0	5.984	-5.984	1.96	Reject
Apr-98	73	2416.0	285.0	5.858	-5.858	1.96	Reject
May-98	73	2556.0	145.0	6.627	-6.627	1.96	Reject
Jun-98	72	2287.0	341.0	5.460	-5.460	1.96	Reject
Jul-98	72	2368.0	260.0	5.892	-5.892	1.96	Reject
Aug-98	72	2368.0	260.0	5.915	-5.915	1.96	Reject
Sep-98	72	2371.0	257.0	5.932	-5.932	1.96	Reject
Oct-98	72	2288.0	340.0	5.466	-5.466	1.96	Reject
Nov-98	72	2287.0	341.0	5.460	-5.460	1.96	Reject
Dec-98	72	2095.0	533.0	4.383	-4.383	1.96	Reject
Jan-99	72	2044.0	584.0	4.097	-4.097	1.96	Reject
Feb-99	72	1815.0	813.0	2.812	-2.812	1.96	Reject
Mar-99	72	1643.0	985.0	1.846	-1.846	1.96	Fail to Reject
Apr-99	72	1530.0	1098.0	1.212	-1.212	1.96	Fail to Reject
May-99	90	2888.0	1207.0	3.382	-3.382	1.96	Reject
Jun-99	94	2866.0	1598.5	2.391	-2.391	1.96	Reject
Jul-99	94	2849.0	1616.0	2.325	-2.325	1.96	Reject
Aug-99	94	3071.0	1394.0	3.162	-3.162	1.96	Reject
Sep-99	94	2817.0	1647.5	2.206	-2.206	1.96	Reject
Oct-99	82	2285.0	1118.0	2.698	-2.698	1.96	Reject
Nov-99	82	2205.0	1198.0	2.328	-2.328	1.96	Reject
Dec-99	82	2125.0	1278.0	1.958	-1.958	1.96	Fail to Reject

Table 4. Wilcoxon Signed-Rank Test, Price Ratio/Exchange Rate Equality (Jan00 – Jun01)

Month	**N**	**T-**	**T+**	**T-***	**T+***	**CV**	**Decision**
Jan-00	81	2152.0	1168.5	2.310	-2.310	1.96	Reject
Feb-00	81	1962.0	1359.0	1.150	-1.150	1.96	Fail to Reject
Mar-00	81	1869.0	1452.0	0.979	-0.979	1.96	Fail to Reject
Apr-00	78	1932.5	1148.5	1.946	-1.946	1.96	Fail to Reject
May-00	81	2319.0	1002.0	3.091	-3.091	1.96	Reject
Jun-00	81	2563.0	758.0	4.236	-4.236	1.96	Reject
Jul-00	81	2093.0	1228.0	2.030	-2.030	1.96	Reject
Aug-00	81	1941.0	1380.0	1.317	-1.317	1.96	Fail to Reject
Sep-00	81	1943.0	1378.0	1.326	-1.326	1.96	Fail to Reject
Oct-00	81	2040.0	1279.0	1.791	-1.791	1.96	Fail to Reject
Nov-00	81	1766.0	1555.0	0.495	-0.495	1.96	Fail to Reject
Dec-00	81	1937.0	1384.0	1.298	-1.298	1.96	Fail to Reject
Jan-01	74	1912.0	863.0	2.816	-2.816	1.96	Reject
Feb-01	79	2269.0	891.0	3.357	-3.357	1.96	Reject
Mar-01	86	2836.0	905.0	4.146	-4.146	1.96	Reject
Apr-01	86	2386.5	1354.5	2.216	-2.216	1.96	Reject
May-01	86	2195.0	1546.0	1.393	-1.393	1.96	Fail to Reject
Jun-01	86	1881.5	1859.5	0.047	-0.047	1.96	Fail to Reject

Table 5. Regression Results for Empirical Version of Equation 2

Dependent Variable:	NEX, Peso/$ nominal exchange rate
Method:	Nonlinear Least Squares
Sample(adjusted):	1997:08 2001:06
Included observations:	47 after adjusting endpoints

Convergence achieved after 8 iterations

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	7.854363	1.559758	5.035629	0.0000
PR	0.282225	0.126836	2.225125	0.0315
IR	0.162112	0.050742	3.194842	0.0027
ICR	-0.390169	0.192680	-2.024957	0.0493
AR(1)	0.674956	0.141979	4.753917	0.0000

R-squared	0.876700	Mean dependent var	9.260426
Adj. R-squared	0.864957	S.D. dependent var	0.624983
S.E. of regression	0.229670	Akaike info criterion	-0.004056
Sum squared resid	2.215434	Schwarz criterion	0.192768
Log likelihood	5.095317	F-statistic	74.65793
Durbin-Watson stat	2.230588	Prob (F-statistic)	0.000000
Inverted AR Roots	0.670001		